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Some International Evidence on Output Stability Under Floating Exchange Rates

by DALLAS S. BATTEN and MACK OTT

With the breakdown of the Bretton Woods system of floating exchange rates in the early 1970s, a new source of short-term price shocks was added to market economies' output-inflation trade-off--namely, those price shocks emanating from exchange rate changes. Many policymakers, businessmen, and columnists have been concerned that governments would attempt to gain competitive advantage in international markets by undertaking monetary policies aimed at lowering their exchange rates. Such a policy, if it worked, would have a double impact on output: (1) the short-run trade-off in domestic goods' markets from accelerating inflation; (2) the short-run competitive advantage in internationally traded goods from an internationally "undervalued" currency.

The output inflation trade-off has been investigated by Lucas (1973a) for a closed economy under fixed exchange rates; it was found that unanticipated price increases were positively associated with increases in output only if the history of price movements was relatively stable. Thus, attempts to exploit the trade-off appeared to be unsuccessful, a conjecture embodied in Friedman's Nobel Lecture (1977). In order to investigate this potential short-run advantage of an undervalued currency, we have adapted Lucas's model and technique to open economies under floating rates. In the first section, an open economy version of the Lucas model is derived along with testable hypotheses concerning exchange rate variability and price movements on real output.

Econometric tests on these hypotheses are reported in section II of the paper, and the closing section sets forth our conclusions and some conjectures to motivate further research.

I. A MODEL OF OUTPUT DISTURBANCES UNDER FLOATING EXCHANGE RATES

In his well-known article (1973a), Lucas derived a simple reduced form model which, although consistent with rational expectations and, therefore, the natural rate hypothesis of unemployment, predicted a short run positive relation between unanticipated price movements and the rate of output. If costly information explains price rigidities (or lags in adjustment) by sellers, we would expect unanticipated price level (or inflation rate) changes to cause changes in output since sellers, in the short run, would be unable to distinguish relative price movements from price level changes. We would also expect, however, that a policy based on this trade-off would fail because if fooling became "systematic" such policy actions would eventually come to be anticipated and, thus, impotent; see Lucas 1973b. Lucas found empirical support for both characteristics in regressions run on annual data for 18 countries from 1954-73, a period at the end of the fixed exchange rate era. In particular, countries with relatively low inflation rates had relatively large positive effects of nominal output on real output; conversely, for the two countries in his sample with high and variable inflation rates, these coefficients were insignificant. As Lucas summarized these results, "...the apparent short-term trade-off is favorable, as long as it remains unused." (Lucas, 1973a, p.333).

Similar to the now discredited premise of an exploitable Phillips Curve is the belief that exchange rate movements might be managed to the fine-tuning-policy antecedents of the Phillips Curve, this motivation suggests an attempt to smooth exchange rate movements which, undamped, might result in the price shocks which advocates of an exploitable tradeoff argued were advantageous. An apt statement of this notion is implicit in Beenstock's summary of "Real Exchange Rate Policy" in Batchelor and Woods (1982):

> "The exchange rate therefore reacts more rapidly to changes in monetary policy than does the domestic price level so that the real exchange rate rises in the short run when monetary policy contracts and rises when monetary policy is expanded. During this period the real exchange rate will be too high or too low. But the blame, so to speak, really rests with the price level and not with the exchange rate. This is because the exchange market is efficient while the internal markets are inefficient. Therefore to blame the exchange rate for being too high when monetary policy is tightened is a miscarriage of justice. For the truth is that domestic wages and prices are too high while the exchange rate is at its appropriate level." (pp. 238-9)

Assuming with Beenstock that foreign exchange markets are efficient and with Lucas that costly information makes domestic prices and wages adjust with a lag, we have a conjectural explanation for the lack of monetary policy independence under floating rates.[2/] That is, central bank policy in the short run may be targeted on suppressing abrupt changes in exchange rates by avoiding relative movements in monetary growth. This can be accomplished by matching changes in the growth rate of the dominant currency determining its exchange rate--the U.S. dollar, for most countries. Over the long run, if monetary growth rates diverge more gradually, the longer time period will also afford, perhaps, a more adequate time span for domestic price and wage adjustments to move in step. Whether such a policy can actually be successful in the sense of lessening fluctuations of domestic real output--hence, employment--or, more succinctly, lowering adjustment costs is not the issue; the question is whether central banks believe this to be feasible and undertake policy on this basis.[3/]

In any case we have, as a first step, to isolate disturbances to domestic production due to unexpected exchange rate movements from disturbances due to unexpected domestic non-traded goods prices. Following Lucas, we have that real output at time t consists of a trend component, y_{nt}, and a cyclic component, y_{ct},

(1) $$y_t = y_{nt} + y_{ct},$$

where

(2) $$y_{nt} = \alpha + \beta t,$$

with all variables expressed in natural logarithms. As in Lucas (1973a, p.327), the cyclic component depends upon deviations of actual prices from expected prices and its own lagged value,

(3) $$y_{ct} = \lambda\, y_{ct-1} + \gamma\, (P_t^{GNP} - E(P_t^{GNP} \mid I_t)),$$

where the GNP deflator is a sectorally weighted average of traded (φ_t) and non-traded goods (P_t) prices:

(4) $$P_t^{GNP} = \varepsilon\, \varphi_t + (1-\varepsilon)\, P_t,$$

(5) $$\varphi_t = P_t^* + e_t^*,$$

where the asterisk indicates the foreign price index and exchange rate with, unless specified otherwise, the foreign country being the United States.[4/]

The deflator expectation in (3) depends upon both non-traded and traded goods' prices given information at the beginning of the period (I_t); thus,

(6) $E(P_t^{GNP} | I_t) = \zeta\, \bar{\varphi}_t + (1-\zeta)\, \bar{P}_t$

where the "-" indicates expected values based on prior information. The sectoral expectation weights are based on the variance of traded and non-traded goods prices and can be expressed approximately as:

(7) $\zeta = \dfrac{Var(P_t)}{Var(\varphi_t) + Var(P_t)}$.

As a final preliminary, our model assumes that relative purchasing power parity (RPPP) holds. Thus, agents in the economies expect exchange rate movements to offset relative price movements between countries. Hence, if purchasing power parity (PPP) holds in t, it will be expected to hold in t+1; if PPP does not hold in t, the discrepancy from PPP in t is anticipated to remain in t+1. Consequently, defining

(8) $\delta_t = \varphi_t - P_t$,

we have

(9) $E(\delta_t | I_t) = \bar{\delta}_t = \delta_{t-1}$.[5/]

Combining (3), (4), (6), (8), and (9) and substituting into (1) yields, finally,

(10) $y_t = y_{nt} + \lambda\, (y_{t-1} - y_{n,\ t-1}) + \gamma\varepsilon\delta_t - \gamma\zeta\bar{\delta}_t + \gamma(P_t - \bar{P}_t)$.

Assuming, once again following Lucas (1973a, p.328), that (10) describes supply behavior based upon correct distributions of expected traded and non-traded goods prices, is equivalent to asserting that the shocks described by (3) result from demand shifts. Denoting GNP by x_t, we have by definition that

(11) $x_t = P_t^{GNP} + y_t$;

further, let Δx_t be a sequence of normal variates with mean g and variance σ_x^2. Combining (11) with (4) and (5) implies

$$x_t = \varepsilon \delta_t + P_t + y_t. \tag{12}$$

The history of the economy then consists of y_{nt} which determines the date, the sequence of demand shocks x_t, x_{t-1}, x_{t-2}, ..., past actual real outputs y_{t-1}, y_{t-2}, ..., and the sequence of international price deviations δ_t, δ_{t-1}, δ_{t-2}, ...

This implies solutions to (12) for P_t and $\bar{P}_t$ of the form

$$P_t = \pi_o + \pi_1 x_t + \pi_2 x_{t-1} + \pi_3 x_{t-2} + \dots + \eta_1 y_{t-1} + \eta_2 y_{t-2} + \dots + \psi_o y_{nt} + \tau_1 \delta_t + \tau_2 \delta_{t-1} + \dots, \tag{13}$$

and

$$\bar{P}_t = \bar{P}_o + \pi_1 (x_{t-1} + g) + \pi_2 x_{t-1} + \dots + \eta_1 y_{t-1} + \eta_2 y_{t-2} + \dots + \psi_o y_{nt} + \tau_1 \delta_{t-1} + \tau_2 \delta_t + \dots \tag{14}$$

To obtain testable hypotheses about the coefficients in (13) and (14) we first solve (12) for y_t and equate this with (10). As Lucas notes (1973a, p.328), this equates supply with demand imposing the market clearing condition on the system. Substituting (13), (14) and (9) for P_t, $\bar{P}_t$, and $\bar{\delta}_t$ in the resulting expression, we obtain, by the method of unknown coefficients, the following reduced form expressions for P_t and y_t:

$$P_t = \frac{\gamma}{1+\gamma} g - \lambda\beta - \lambda y_{t-1} - (1-\lambda) y_{nt} + \frac{1}{1+\gamma} x_t + \frac{\gamma}{1+\gamma} x_{t-1} - \varepsilon \delta_t + \gamma(\zeta-\varepsilon)\delta_{t-1}, \tag{15}$$

(16) $$y_t = -\frac{\gamma}{1+\gamma} g + \lambda\beta + \lambda\, y_{t-1} + (1-\lambda) y_{nt}$$

$$+ \frac{\gamma}{1+\gamma} \Delta x_t + (\varepsilon-\varepsilon)\delta_t - \gamma(\zeta-\varepsilon)\delta_{t-1},$$

The reduced form estimating equations for inflation and cyclic real output deviations are then obtained, respectively, by taking the first difference of (15) and by applying (3) to (16):

(17) $$\Delta P_t = -\beta - \lambda \Delta y_{c,t-1} + \frac{1}{1+\gamma} \Delta x_t + \frac{\gamma}{1+\gamma} \Delta x_{t-1}$$

$$- \varepsilon\, \Delta\delta_t + \gamma(\zeta-\varepsilon)\Delta\delta_{t-1},$$

(18) $$y_{ct} = -\frac{\gamma}{1+\gamma} g + \lambda y_{c,t-1} + \frac{\gamma}{1+\gamma} \Delta x_t + (\varepsilon-\varepsilon)\delta_t - \gamma(\zeta-\varepsilon)\delta_{t-1}.$$

The implications of cyclic deviations of real output and unanticipated GNP shocks on the inflation rate and real output are, of course, the same as those derived by Lucas. That is, in (17), the change in GNP has an effect on inflation (over a two-quarter span) only to the extent that it exceeds trend real output growth (measured by β) and the lagged impact of the cyclic real output deviation. Similarly, in (18), nominal demand shifts, Δx_t, cause real output to deviate from trend only if they differ from the trend growth of demand (g). However, (17) and (18) also contain implications for inflation and real output due to changes in the discrepancy from PPP and the difference, if any, between traded goods share in nominal GNP (ε) and the effect of world prices on the domestic GNP deflator.

If RPPP holds there will be no change in the PPP discrepancy from period to period; thus, (17) implies under RPPP that there would be no impact of sustained deviations of international prices from PPP on domestic inflation. Moreover, even when δ_t does change -- by assumption, unexpectedly according to (9) -- (17) implies that there will be a contemporaneous negative effect on inflation, proportional to the traded goods share in GNP, and the lagged impact will be non-zero proportional to difference between the sectoral expectation weight ζ and the traded goods share ε.[6/] In contrast, (18) asserts that real output is negatively affected only by lagged δ in proportion to the difference between the sectoral weight and the traded goods share; the contemporaneous impact of δ works through Δx. Somewhat puzzling, then, is the implication from (18) that a sustained departure from PPP can have real effects, if the expectation formation weight of the international price differs from its trade share in the national income accounting identity.

Finally, there are several additional implications of the model summarized in (17) and (18):

a) the intercept in (18) is negative and equal to the product of coefficient on Δx_t and trend growth rate of x_t;

b) the coefficient of δ_{t-1} in (18) is equal to the difference of ζ (from (7)) and ε times γ implied by coefficient of Δx_t;

c) an estimated autoregression of δ_t will be consistent with RPPP--(9) and footnote (5);

II. EMPIRICAL RESULTS

Equations (18) and (1') (footnote 5) were estimated for seven developed countries which reflect a wide range of financial institutions and central bank behavior: Canada, France, Germany, Italy, Japan, the

Netherlands and the United Kingdom. To account for lagged variables, the estimation period spans IV/1973 to the most recent period for which data are available (which ranges from IV/1982 to II/1983). The variables employed or the ones constructed should be evident from the preceding discussion.

Because one of the basic assumptions of the model--that relative purchasing power parity holds--is testable, we investigate this first by estimating equation (1') and testing the hypotheses enumerated in footnote 5. These results are presented in table 1. They indicate that, while the hypotheses that $\theta_0 = 0$ and $\theta_1 = 1$ cannot be rejected for any country, the hypothesis that the residuals are white noise is rejected for Canada, France, Italy and Japan. That is, the implications of RPPP are not supported for these countries. When an additional lagged value of ϕ (ϕ_{t-2}) is added to (1'), however, the residuals are white noise. Consequently, it appears that the assumption that RPPP holds, at least approximately, is not completely unfounded.

The results from estimating equation (18) are reported in table 2. From inspection of the Durbin h, one notices that this equation, when estimated for Canada, France, Italy, and the U.K., exhibits significant first-order autocorrelation. The estimations correcting for this are reported in table 3. Note the relatively small differences between the equation as initially estimated and the one adjusted for autocorrelation--a fact that lends additional evidence to the specification of (18). Furthermore, as predicted, the constant terms are negative and the estimated coefficients of Δx_t and $y_{c,t-1}$ lie between zero and one.[7/]

Another way of evaluating the estimates of (18) reported in Tables 2 and 3 is by comparing the relative responsiveness of real output--measured by the coefficient of Δx_t--with the relative variability of the inflation rate--measured by the variance of the quarterly inflation rate. As mentioned earlier Lucas (1973a, pp. 332-333) found that the two countries in his study with high and variable inflation rates (relative to the other 16) had insignificant responses to nominal GNP changes. The rationale for the ordering is that a higher variability of inflation imputes less information about real output changes from a given Δx_t. As shown in table 4, four of the countries in our sample had inflation variability of half or less that of the other three countries--Germany, France, Canada and the Netherlands had comparatively low variances of inflation. These low inflation variance countries, as expected, also had appreciably higher coefficients on Δx_t.[8/] Also, as indicated above, the negative of the ratio of the constant and the coefficient of Δx_t should approximately equal the trend rate of growth of x. This is generally supported by the first two columns of table 5.[9/]

The results in tables 2 and 3 concerning the coefficients of δ_t and δ_{t-1} provide some general support for the model. Except for Canada, the coefficients of δ_t are neither sizable absolutely nor statistically significant, as hypothesized. The coefficients of δ_{t-1}, however, conform less to the model-generated hypotheses. In particular, their signs should be the same as $(\zeta-\varepsilon)$--a condition satisfied only by Canada, Japan and the U.K. (note last two columns of table 5). Furthermore, only the coefficients for Canada, France and the

Netherlands are statistically significant. Consequently, one might conclude that, in general, unexpected price changes with foreign origins have little influence on domestic output. This finding might not be two surprising once the calculated ζs are examined (table 6); these indicate that, except for Canada, Japan, and the U.K., the relative weight given to foreign prices in determining the expected general price is very small. As a result, one might expect foreign price movements to have relatively little explanatory power.

III. CONCLUSIONS AND CONJECTURES

The motivation for this paper was the empirical observation (Batten-Ott, [1983]) that monetary growth rates of the industrialized economies have not been independent under floating exchange rates; there has been a significant short-run dependence of foreign monetary growth rates on U.S. money growth, but no long-run dependence. A first step in explaining this short-run interdependence was to investigate the possibility that exchange rate movements cause real output fluctuations. If so, we conjectured, this might induce monetary authorities to undertake policies to avoid abrupt movements in exchange rates but allow persistent exchange rate movements; this would result in the observed monetary growth relation.

In order to test this smoothing motivation for policy interdependence, we extended the Lucas (1973a) model to the open economy setting implied by floating exchange rates. Unfortunately, while the model answers the questions addressed to it--namely, deviations from PPP have not generally caused real output deviations--we do not have a definitive answer to our monetary policy puzzle. That is, the lack of significance of δ_t in (18) may be interpreted either as showing that floating exchange rate deviations from PPP do not cause real output fluctuations, or, alternatively, that central bank policies have smoothed such deviations to an extent that has obviated their impact. Thus, in order to address this issue we will have to generate a different model or augment this one so that the policy linkage is articulated explicitly.

While the estimates do not illuminate the monetary policy issue, they do conform to the model's predictions for the domestic variables' coefficients--that is, GNP growth, output responses to GNP changes, and the lagged cyclic output coefficient. Moreover, with the exception of Canada, δ_t is not an explanation for cyclic output fluctuations. The model's estimates do not seem to be affected by serially correlated residuals; as a comparison of Tables 2 and 3 reveals, the coefficients change only slightly when the estimated equations are corrected for serial correlation.

The aspect of the model that is most puzzling is the coefficient on δ_{t-1}. First, there is the theoretical puzzle as to why there should be a sustained impact on real output fluctuations from a lagged price, particularly in a rational expectations model. Equation (18) implies that the difference, if any, would be due to the discrepancy between the expectational weight of traded goods prices in the deflator, ζ, and the national income accounts weight of traded goods in the deflator, ε. This difference is, by computation, generally substantial, as shown in Table 4, yet it seems inconsistent with relatively efficient markets (and approximate RPPP) that lagged δ should affect current output. The second part of the puzzle is, of course, that empirically, δ_{t-1} does not, except for Canada, have a great impact; the coefficient is also significant, but small, for France and the Netherlands.

Given these preliminary findings, our current work is focused on two fronts. From the modeling standpoint, we are attempting to resolve the theoretical puzzle of the coefficient of δ_{t-1} and trying to incorporate explicitly the monetary policy stance of the central bank. From the data side, we are constructing price series for non-traded and traded goods which better capture this distinction than the CPI services and PPI series used in this paper. In addition we are considering the question of the relation of energy price movements to both real output fluctuations and exchange rates; note that three of our sample economies are net energy exporters (Canada, the Netherlands, and the U.K.) while the other four are nearly wholly dependent on imported energy.

In summary, this relatively simple extension of the Lucas model to open economies under floating exchange rates performs reasonably well in these preliminary results. In particular, floating exchange rates *per se* do not appear to be a cause of output fluctuations. Consequently, one might conclude that, in general, unexpected price changes with foreign origins have little influence on domestic output. Whether floating rates can be blamed for output fluctuations in the model when monetary policy and energy price movements are accounted for and when we have better price indexes for traded and non-traded goods remains to be seen.

FOOTNOTES

1/ It is not necessary to explain all such correlations between monetary growth rates under floating rates since, as Johnson (1969) argued there are advantages to a fixed rate for small economies:

> "...under a flexible rate system most countries would probably peg their currencies to one or another major currency, so that much international trade and investment would in fact be conducted under fixed rate conditions, and uncertainty would attach only to changes in the exchange rates among a few major currencies or currency blocs (most probably, a U.S. dollar bloc, a European bloc, and sterling, though in the event sterling might be included in one of the other blocs)." (Johnson, pp. 12-13)

2/ Note that this is subtly, but significantly, different from Beenstock's characterization of inefficient domestic markets. Costly information causing price adjustment to lag real or financial changes is inconsistent with neither efficient markets nor rational expectations of individuals in those markets.

3/ Note, given the absence of market-determined compensation to central bankers (and their overseeing governments), that there is less justification for assuming optimality in their decisions than in private markets.

4/ In this sense, under floating exchange rates, Lucas's equation (3) (p.327, 1973a) would be a misspecification for an open economy. Furthermore, it could be argued that, for members of the European Monetary System, the "foreign" country should be Germany. We did incorporate this possibility in our empirical analysis, but found the results to be generally inconsistent with the model, and consequently, they were not reported.

5/ More concisely, this formulation of RPPP asserts that the estimation of

$$(1') \qquad \delta_t = \theta_o + \theta_1 \delta_{t-1} + \mu_t$$

will result in θ_1 not significantly different from unity, θ_0 not significantly different from zero and μ_t will be white noise.

6/ As Lucas subsequently observed (1976) both the inflation and real output deviations equations are obtained from the same identify--(12) in our case. Hence, the equations for inflation and real output deviations (17) and (18), are not independent. Since the two equations do led to alternative estimates of combinations of the structural parameters, however, we intended initially to investigate both relationships using the CPI for services as a proxy for non-traded goods' prices in equation (17). The results were generally inconsistent with the model, and so we have left this task for subsequent research, choosing to focus our attention on the estimation of equation (18).

7/ The only apparent exception to this is the coefficient on Δx_t for the Netherlands; however, the point estimate of 1.04 is not significantly different from 1.0, which, by an application of L'Hospital's Rule, can be shown to be the limit of $\gamma/(1+\gamma)$ as γ becomes indefinitely large. The Netherlands is the most open economy in our sample--its ϵ is .55 which is twice as large as the next most open economies, Germany and the U.K. Thus, it is reasonable that the economy whose domestic production is most subject to traded goods competition should also have the highest implied supply elasticity.

8/ Japan's inflation variability is especially noteworthy since its average inflation rate was low, second only to that of Germany. The Japanese inflation rate was very high in 1973--24.3 percent annual rate--but relatively stable and low from 1962-1971--5.6 percent--and from 1977 to 1982--just over 4.5 percent. Thus, it is the impact on expectations of the era of oil price shocks--1971-1977, mean rate 12.9 percent, and 1980, 8.0 percent--which appears to have raised the inflation variance and lowered the impact on output of changes in nominal GNP.

9/ Whether the differences between columns 1 and 2 in table 4 are statistically significant is not easily tested since determining the distribution of the ratio of two normal random variables (which are not independent) is not a trivial matter.

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Table 1
Investigation of RPPP

Country	$\theta_0 = 0$ 1/	$\theta_1 = 1$	Durbin h
Canada	Yes	Yes	2.93*
France	Yes	Yes	2.96*
Germany	Yes	Yes	1.65
Italy	Yes	Yes	2.31*
Japan	Yes	Yes	3.13*
Netherlands	Yes	Yes	1.92
U.K.	Yes	Yes	1.74

*Statistically significant at 5 percent level.

1/ "Yes" indicates cannot reject hypothesis that $\theta_0 = 0$ at the 5 percent level.

2/ "Yes" indicates cannot reject hypothesis that $\theta_1 = 1$ at the 5 percent level.

Table 2
Estimation of Equation 18

	Estimated Coefficient of						
Country	Constant	Δx_t	$y_{c,t-1}$	δ_t	δ_{t-1}	$\bar{R}^2$	DW/h
Canada	-.023*	.763*	.952*	.182*	-.193*	.95	1.21/2.43*
	(9.16)	(9.63)	(23.86)	(3.39)	(3.66)		
France	-.020*	.616*	.973*	.032	-.034	.88	1.15/2.97*
	(4.82)	(5.83)	(12.59)	(1.45)	(1.54)		
Germany	-.013*	.884*	.889*	.007	-.022	.95	1.76/0.65
	(4.80)	(8.17)	(16.60)	(0.38)	(1.23)		
Italy	-.022*	.573*	.811*	-.0003	-.024	.90	1.36/2.13*
	(5.01)	(8.92)	(13.70)	(0.01)	(0.72)		
Japan	-.006*	.238*	.668*	.018	-.048	.76	1.58/1.48
	(2.41)	(2.51)	(9.35)	(0.73)	(1.86)		
Netherlands	-.016*	1.04*	.899*	.026	-.076*	.90	2.74/1.30
	(6.17)	(14.46)	(11.20)	(0.97)	(3.05)		
U.K.	-.015*	.358*	.838*	-.005	.010	.76	1.37/2.25*
	(2.77)	(2.80)	(8.40)	(0.10)	(0.21)		

Absolute value of t-statistics in parentheses.
*Statistically significant at the 5 percent level.

Table 3
Estimation of Equation (18) Adjusted for First-Order Autocorrelation

Country	Constant	Estimated Coefficient of Δx_t	$y_{c,t-1}$	δ_t	δ_{t-1}	ρ	DW/h
Canada	-.020*	.673*	.907*	.154*	-.153*	.49*	1.91/0.25
	(7.33)	(9.37)	(17.53)	(3.13)	(3.08)	(3.51)	
France	-.018*	.606*	.789*	-.046	-.018*	.61*	2.37/1.51
	(4.08)	(7.60)	(7.81)	(1.84)	(2.35)	(4.75)	
Italy	-.015*	.460*	.731*	.004	-.036	.60*	1.92/0.24
	(2.49)	(8.64)	(9.58)	(0.15)	(1.20)	(4.62)	
U.K.	-.025*	.543*	.383*	-.031	.010	.85*	1.97/0.11
	(2.12)	(7.11)	(3.29)	(1.05)	(0.34)	(9.95)	

Absolute value of t-statistics in parentheses.
*Statistically significant at the 5 percent level.

Table 4
Responsiveness of Real Output to Changes in Demand Compared with Relative Inflation Variability

	Response		Inflation[2]		
Country	Coefficient of Δx_t	Rank	Mean	Variance	Rank[3]
Canada	.673[1]	3	.024	$.069 \times 10^{-3}$	3
France	.606[1]	4	.026	$.053 \times 10^{-3}$	2
Germany	.884	2	.011	$.031 \times 10^{-3}$	1
Italy	.460[1]	6	.040	$.163 \times 10^{-3}$	5
Japan	.238	7	.015	$.236 \times 10^{-3}$	6
Netherlands	1.04	1	.017	$.092 \times 10^{-3}$	4
U.K.	.543[1]	5	.033	$.294 \times 10^{-3}$	7

1/ Estimate from equation corrected for first-order autocorrelation--table 3.
2/ Quarterly rates of inflation and variances.
3/ Ranking by lowest variance of inflation rate.

Table 5
Comparison of Estimated Parameters Implied by Regression Coefficients with their Computed Counterparts, 1973IV-1983

	Trend Growth Rate of GNP: g		Mean Difference Between Proportional Non-traded Goods Price Variance and Traded Goods Share: $\zeta-\varepsilon$	
Country	Estimated 1/	Computed 2/	Estimated 3/	Computed
Canada	.030	.030	.074	.091
France	.030	.032	.012	-.187
Germany	.015	.015	.003	-.276
Italy	.033	.046	.042	-.140
Japan	.025	.024	.154	.143
Netherlands 4/	.015	.012	--	-.485
U.K.	.046	.035	-.008	-.109

1/ Constant term in estimate of (18) divided by coefficient of Δx_t in estimate of 18--see Table 2.
2/ Quarterly growth rate.
3/ Coefficient of δ_{t-1} in estimate of (18) divided by γ as inferred from coefficient of Δx_t in estimate of (18)--see Table 2.
4/ Estimated $\zeta-\varepsilon$ not defined since estimated γ is infinite; see note 7.

Table 6
Calculated ξ

Country	ξ
Canada	.352
France	.015
Germany	.010
Italy	.088
Japan	.284
Netherlands	.070
U.K.	.168